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Does an expanding low-pay sector decrease structural unemployment? Evidence from Germany

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Does an expanding low-pay sector decrease structural unemployment? Evidence from Germany

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Abstract

Low-pay work has been increasing in prevalence in many industrial countries. Following standard wage/price-setting theory, this increase should reduce structural unemployment, because labour market flexibility increases and labour costs decrease. However, a Keynesian perspective challenges this claim, if the associated increase in investment demand does not sufficiently compensate for the negative effects on consumption. This research empirically investigates the theoretically uncertain impact of the relationship between the extensiveness of the low-pay sector and structural unemployment. Data from Germany, where the expansion of the low-pay sector has been declared the goal of the labour market policy, during the period from 1991 to 2008, indicate a positive impact of the growing low-wage sector on structural unemployment. Moreover, some indications suggest an opposite direction of causality, such that changes in the level of structural unemployment affect the share of low-wage earners. This effect is asymmetrical with respect to positive and negative changes, which seems to reflect downward wage rigidity, as caused by labour market institutions.

NAIRU; low-wage; asymmetric error correction; cointegration; VECM

JEL Classifications: C32, E24, J31

1 Introduction

As in many other OECD countries, the German low-wage sector has been growing for several years, partially as an intended result of labour market reforms. For example, reforms enacted in the course of the so-called Hartz legislation are based on the argument that high levels of long-run unemployment are caused by institutional and structural factors, such as labour market rigidities and an overly even wage structure. Krugman (1994) argues that low levels of unemployment are permanently possible, at least at the price of higher wage inequality: The comparison to Anglo-Saxon economies shows that inequality and unemployment 'are two sides of the same coin'. According to this hypothesis and standard wage/price-setting theory, an expanding low-wage sector should reduce structural unemployment, because it creates additional labour demand by decreasing labour costs, especially with respect to less-skilled work. Furthermore, it increases labour market flexibility, in that low-pay jobs tend to be related to fixed-term, marginal or temporary relationships, which are usually less rigid than standard labour contracts (Kalina and Weinkopf (2008)). However, if a growing low-pay sector is accompanied by decreases in consumption, it could exert multiplier and accelerator effects and thereby lead to a downturn in aggregate demand. Such demand drops likely influence the supply side, with negative effects on employment. Obviously, the expected impact of an expanding low-wage sector on structural unemployment thus is not clear-cut; the anticipated outcome depends mainly on the theoretical approach chosen. To attempt to resolve this question, we investigate the relationship between both variables empirically.

In this approach, our research relates to literature on the impact of institutions on (structural) unemployment (e.g. Bassanini and Duval (2006); Blanchard and Wolfers (2000); Gianella et al (2008); Nickell et al (2005)). Prior research has tested several structural and institutional variables, such as productivity, terms of trade, interest rates, unemployment benefits, tax wedges, employment protection legislation, union coverage, or expenditures on active labour market policies. We contribute to this strand of literature, because the low-pay sector variable has not yet been considered in this context, though it reflects institutional factors and thus labour market flexibility. In Germany, two institutional factors have been contributing to the expansion of low-pay work. First, the German system of industrial relations has changed. Union coverage declined, because the number of companies belonging to employers' associations decreased, as did the share of employees with union membership. Thus, the number of collective agreements declared generally binding has decreased (Bosch and Kalina (2008)). Second, the labour market has been deregulated (Carlin and Soskice (2009)), which has encouraged more part-time, fixed-term, marginal, and temporary employment, as well as tightened sanctions for unemployed workers who reject job offers (Kommission (2002)).[1]

The relationship between low-pay work and unemployment has been recently addressed. Garz (2010) uses single-equation results to indicate a slightly positive impact of the extent of the low-wage sector on structural unemployment. We extend that study by allowing for interdependencies among the variables; specifically, we employ a vector error correction model (VECM) and do not rule out the reverse direction of causality a priori. In addition, we estimate an asymmetric error correction model (AECM), because the study variables may react to positive and negative shocks in non-linear ways. Economic policy and (labour market) regulation may exert positive changes in the variables that have different magnitudes than the negative changes.

[1]The extent of the low-wage sector captures the share of employees at the lower part of the wage distribution. In addition to institutional factors, the low-wage variable is therefore affected by factors that generally influence the degree of wage inequality (e.g. skill-biased shifts in labour supply due to educational expansion and migration movements, increased female labour force participation rates [Bosch and Kalina (2008)], skill-biased technological change [Krueger (1993); Acemoglu (1999)], and the effects of global trade liberalisation [Freeman (1995); Fitzenberger (1999)]).

In the next section, we elaborate on the theoretical background for our study. Thereafter, we explain our econometric strategy and describe the data. We present and discuss the results from the VECM and AECM estimations before we conclude with some implications in the last section.

2 Theoretical background

2.1 Standard wage/price-setting theory

Structural unemployment is an inherently unobserved variable. In standard wage/price-setting theory, it is defined as the unemployment rate that prevails when the price- and wage-setting functions are in equilibrium—often referred to as the non-accelerating inflation rate of unemployment (NAIRU). In equilibrium then, all wage and price expectations are fulfilled, and temporary shocks are absent. The theoretical impact of the extent of the low-wage sector on the NAIRU can be evaluated according to price- and wage-setting processes. For example Layard et al (1991) (361-396) propose setting prices under incomplete competition as a constant mark-up over labour costs. Inter alia, labour costs are affected by short- and long-run supply-side factors, such as oil or import price shocks, capital costs and the intensity of competition. Wage setting then is an outcome of the collective bargaining process, which depends on short- and long-run factors that influence bargaining power, such as changes in the trade balance, unemployment benefits or union coverage. In this regard, the extent of the low-wage sector does not affect price directly but rather influences wage setting. An growing low-pay sector implies a reduction in real labour costs and thus increases labour demand, whereas a decreasing low-wage sector has the opposite effect. These changes relate to long-run institutional or structural factors that determine labour demand and supply in the lower wage distribution.

The long-run path of the NAIRU only depends on institutional and structural factors though, because price and wage surprises and temporary shocks merely affect the Phillips curve trade off, which is assumed to exist only in the short run. A lasting low-wage expansion leads to a decline in the NAIRU, as we depict in Figure 1. Due to reduced real labour costs, the wage-setting curve (WSC) shifts to the right, whereas the price-setting curve (PSC) remains unchanged. The economy moves from point A to its new equilibrium B. If the PSC is downward sloping, as in the graph, real wages fall from w^* to w_1^*. More important though, the NAIRU decreases from u^* to u_1^*.

2.2 Keynesian demand effects

The demand side of the economy plays only a passive role in the NAIRU model. That is, employment appears determined only by the supply side, and aggregate demand simply adjusts to a certain employment equilibrium. The NAIRU then is exogenous with respect to the demand side. Institutional factors—in this context, the extent of the low-pay sector—are supposed to affect the long-run position of the NAIRU only through the supply side, whereas demand effects may have at best temporary repercussions for

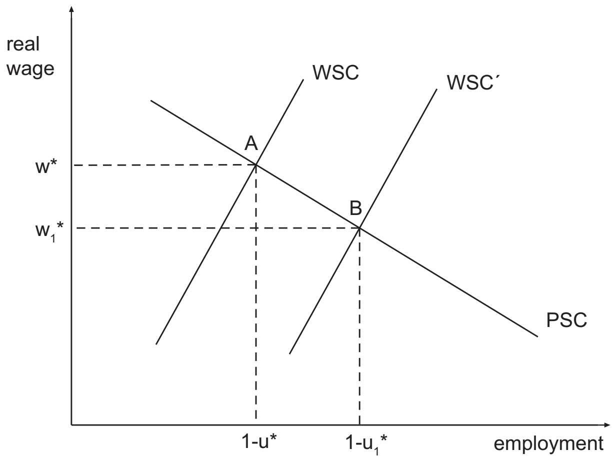

Figure 1: Low-wage expansion and the NAIRU

unemployment. However, both theoretical and empirical issues challenge this notion. From a Keynesian point of view, employment is determined by aggregate supply and demand outcomes in goods markets. Thus, malfunctions in the labour market explain only a small fraction of total unemployment. Empirical studies support this argument, as summarised by Blanchard and Wolfers (2000): 'labour market institutions do not appear to explain the general evolution of unemployment over time.'

Some authors have incorporated features that allow the NAIRU to be endogenously determined by demand factors. For example, some models account for hysteresis effects, which would imply permanent effects of shocks on employment through the demand side, such as those caused by a restrictive monetary policy (Ball (1997), Ball (1999)). Others have emphasised the role of capital accumulation (Blanchard et al (1997); Arestis et al (2007); Stockhammer and Klaer (2011)). Capital accumulation may affect employment through two channels: First, low investment demand, amplified by a standard multiplier mechanism, results in low aggregate production and unemployment. This scenario may exert hysteresis effects on the labour market and permanently lower capital stock accumulation, which initiates—through the mechanism of poor profitability expectations—a self-energising period of low investment. Second, Rowthorn (1999) has shown that even in the NAIRU framework, capital accumulation affects employment if the elasticity of substitution between capital and labour is less than 1. In this case, an increase in capital is accompanied by an increase in the share of labour.

Because the extent of the low-wage sector affects the level of real wages (i.e. a high share of low-wage earners implies a lower real-wage level than does a low share, ceteris paribus), the relationship between the level of real wages and aggregate demand is crucial for evaluating the theoretical effect of an expanding low-wage sector on (structural) unemployment. In a framework where employment is determined by goods rather than the labour market, wages have a cost as well as a demand effect (Bhaduri and Marglin

(1990)). The overall effect then depends on the prevailing goods market regime.[2] In a profit-led regime, an expanding low-wage sector results in lower real unit labour costs. The negative impact on consumption due to lower wage income is more than compensated for by higher investment demand, which leads to an overall increase in aggregate demand and production, as well as in employment. In contrast, if the prevailing regime is wage-led, investment demand is inelastic with respect to changes in real wages. If the increase in investment demand is not sufficiently high to compensate for the negative consumption effect, it results in less aggregate demand, production and employment.[3] Thus, the impact of a low-wage expansion depends on the goods market regime.

3 Econometric strategy

3.1 The vector error correction model

Both theoretical approaches acknowledge that the extent of the low-pay sector can causally influence the level of structural unemployment. To test whether there are repercussions from structural unemployment to the low-pay sector, we allow for interdependencies and employ a VECM. The structural form for the determination of the $m \times 1$ vector of variables z_t is given by:

$$A\Delta z_t = \tilde{a} + \tilde{b}t - \tilde{\Pi} z_{t-1} + \sum_{i=1}^{p-1} \tilde{\Gamma}_i \Delta z_{t-i} + \epsilon_t \; . \tag{1}$$

Matrix A contains contemporaneous structural coefficients, whereas $\tilde{\Pi}$ and $\tilde{\Gamma}$ contain dynamic coefficients that relate Δz_t to past values of z_t. All matrices are of size $m \times m$, whereas $\tilde{a}$ and $\tilde{b}$ are $m \times 1$ vectors of structural coefficients. The $m \times 1$ vector of structural disturbances ϵ_t is serially uncorrelated with zero means and a positive definite variance covariance matrix, Ω.

If there are $0 < r < m$ cointegrating vectors, then β will be an $m \times r$ matrix, and $\tilde{\Pi}$ will be of rank r:

$$\tilde{\Pi} = \tilde{\alpha}\beta' \; , \tag{2}$$

where the linear combination $\beta' z_t$ is I(0) and refers to the deviation from equilibrium, such that the matrix $\tilde{\alpha}$ $(m \times r)$ captures the adjustment coefficients. Pre-multiplying equation (1) by A^{-1} yields the reduced form VECM:

$$\Delta z_t = a + bt - \Pi z_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta z_{t-i} + v_t \; , \tag{3}$$

[2] Bowles and Boyer (1995) find evidence that the German economy follows a profit-led demand regime, whereas Naastepad and Storm (2007) and Hein and Vogel (2008) find evidence for a wage-led one.

[3] Carlin and Soskice (2009) argue that an increase in labour market flexibility strengthens households' precautionary savings motive 'in response to the concerns about cutbacks in the welfare state generated by government policy', which could amplify the negative income effect.

where $a = A^{-1}\tilde{a}$, $b = A^{-1}\tilde{b}$, and $\Gamma_i = A^{-1}\tilde{\Gamma_i}$. In addition, $\Pi = A^{-1}\tilde{\Pi} = A^{-1}\tilde{\alpha}\beta' = \alpha\beta'$, where $\alpha = A^{-1}\tilde{\alpha}$ and $v_t = A^{-1}\epsilon_t$ are the reduced-form errors. The corresponding variance covariance matrix is Σ, with $\Omega = A\Sigma A'$ (Garratt et al (2006)).

3.2 The asymmetric error correction model

The standard VECM assumes that a response of a variable to an impulse from another variable is symmetric: Negative changes should have the same magnitude as positive ones. In contrast, the AECM is not restricted to this assumption but allows for asymmetric responses. We start with the following asymmetric (cointegrating) long-run regression (for details, see Shin et al (2009)):

$$y_t = \beta^{+\prime} x_t^+ + \beta^{+\prime} x_t^- + u_t \,, \tag{4}$$

where y_t, x_t^+ and x_t^- are I(1) variables, and the last two series are the positive and negative partial sum processes, defined as:

$$x_t^+ = \sum_{j=1}^{t} \Delta x_j^+ = \sum_{j=1}^{t} max(\Delta x_j, 0) \,, \tag{5}$$

$$x_t^- = \sum_{j=1}^{t} \Delta x_j^- = \sum_{j=1}^{t} min(\Delta x_j, 0) \,. \tag{6}$$

The coefficients β^+ and β^- refer to the asymmetric long-run impacts. Equation (4) can be rewritten as an autoregressive distributed lag (ARDL) representation and further transformed into the AECM:

$$\Delta y_t = \rho y_{t-1} + \theta^+ x_{t-1}^+ + \theta^- x_{t-1}^- + \sum_{j=1}^{p-1} \varphi_j \Delta y_{t-j} + \sum_{j=0}^{q} (\pi_j^+ \Delta x_{t-j}^+ + \pi_j^- \Delta x_{t-j}^-) + \epsilon_t \,. \tag{7}$$

The asymmetric long-run coefficients are computed as $\widehat{\beta^+} = -\widehat{\theta^+}/\widehat{\rho}$ and $\widehat{\beta^-} = -\widehat{\theta^-}/\widehat{\rho}$. Equation (7) then can be estimated using ordinary least squares (OLS). In this context, we are particularly interested in the following tests:

- A long-run level relationship between the variables. We apply the bounds-testing approach proposed by Pesaran et al (2001) (PSS) and Shin et al (2009), to test the null hypothesis of no level relationship between the variables versus the existence of such a long-run relationship, using the PSS F-statistics:

  $$H_0\colon \rho = \theta^+ = \theta^- = 0 \,,$$

  where ρ, θ^+ and θ^- are the relevant coefficients in equation (7).

- Long-run symmetry. We test the null hypothesis of unity for the long-run parameters, β^+ and β^-, against the alternative hypothesis of diversity, using Wald statistics:[4]

$$H_0 \colon \beta^+ = \beta^- \, .$$

4 Data

Because structural unemployment is not observable, we use the Kalman filtered unemployment rate (time-varying NAIRU) as a proxy. The extent of the low-wage sector (LWS) is calculated with data from the German Socio-Economic Panel (SOEP), version 2008, of the German Institute for Economic Research (DIW). We provide the details of the NAIRU estimation and the calculation of the low-pay extent in the Appendix.[5] As control variables, we use the real interest rate (RLTIR), the vacancy ratio (VACRATIO) and the log of the oil price (OILPRICE). The real interest rate refers to the nominal long-term interest rate deflated by the gross domestic product (GDP) deflator; it provides a proxy for monetary factors that affect the economy. The vacancy ratio is defined as the ratio of vacancies to the number of unemployed, which indicates mismatch unemployment. The oil price aims to capture exogenous, unexpected supply shocks.[6] We restrict our choice of controls to these three variables to keep the identification of our econometric models feasible. All data are available from at least 1984Q3 to 2008Q4. However, stability issues related to the time before German reunification in some models prompted us to restrict the analysis to the period from 1991Q1 to 2008Q4.

We plot the data in Figure 2. After the reunification shock, the NAIRU initially tended to increase, from about 8.9% to 9.3% in 1998, at which point it began falling to its lowest post-reunification value of about 8.8%. Thereafter, the NAIRU rose again until it reached its peak of approximately 9.7% in 2005, then began another decline. The second time-series of primary interest, the extent of the low-wage sector, has its lowest values in the years immediately following German reunification. From 1995 onward, it has shown a clear upward trend, implying an increase of low-wage earners from about 16% in 1995 to 22% in 2006. The visual inspection of the data thus offers a first indication that the relationship between the variables may not be negative as predicted by standard wage/price-setting theory but rather should be characterised as ambiguous, if not positive.

[4]We also allow for short-run asymmetry but do not test it explicitly.

[5]We do not attempt to incorporate the low-wage extent or other institutional variables already in the Kalman filter procedure. When applying the filter, the NAIRU is treated as a variable, determined solely stochastically on the basis of actual unemployment in a Phillips curve context. The stochastic process is assumed to capture all influences affecting the NAIRU, so an omitted variable bias cannot occur.

[6]The data for the vacancy ratio and the real interest rate are seasonally adjusted and provided by the OECD. The oil price data are obtained from the FRED database of the Federal Reserve Bank of St. Louis.

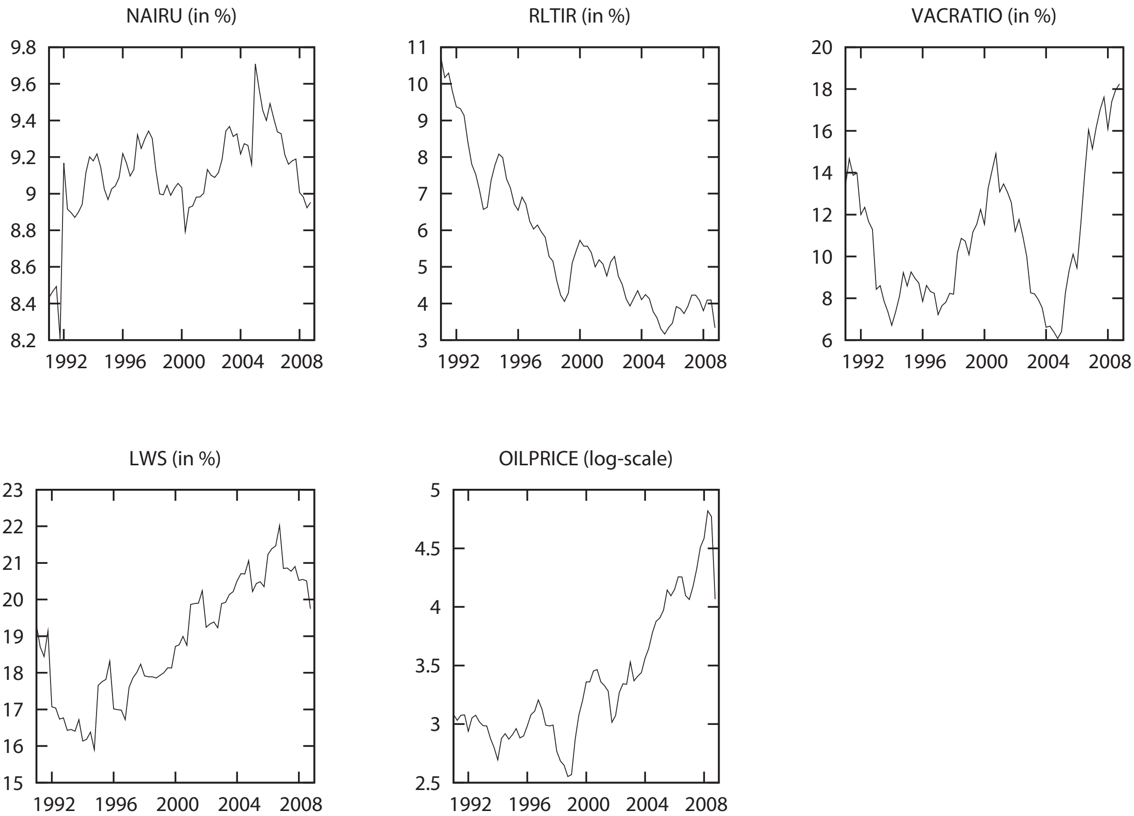

Figure 2: Time-series plots of the data

Before we conduct our cointegration analysis, we evaluated the order of integration of each time-series. The insight from usual stationarity tests is ambiguous (see Tables 5 and 6 in Appendix for details): The ADF-GLS test suggests that the vacancy ratio might be I(0) and the low-wage share and the oil price even I(2). In contrast, the KPSS test suggests that all variables are I(1), even if the null hypothesis of stationarity in the level of the vacancy ratio can only be rejected at the 10% level. Given the lack of power of stationarity tests and considering our small sample size, we assume all variables to be I(1).

5 Results

5.1 Vector error correction estimates

We estimate four models. The first comprises the following variables: low-wage share, NAIRU, real interest rate and vacancy ratio. Models 2, 3 and 3R also incorporate an unrestricted impulse dummy, which takes a value of 1 at date 2001Q1 and 0 at other times to account for an outlier. Model 3 includes the oil price as a weakly exogenous[7] I(1) series, as well as its changes as unrestricted variables. In model 3R (restricted model 3), we test the hypothesis that the NAIRU does not enter the long-run equilibrium relationship by setting the corresponding alpha and beta coefficients to 0.

According to different information criteria, a VECM(5) is appropriate for all three models, irrespective of the deterministics (i.e. with constant, with constant and trend, with and without dummy variables and the level of the oil price). Johansen trace tests suggest, for each of the three models, one long-run level relationship (see Tables 7 and 8).

Accordingly, we estimate the four VECMs with unrestricted intercepts and normalise the low-wage share to the unit level.[8] The estimation and corresponding test results are summarised in Table 1. The fit of the LWS and NAIRU equations are much better when the impulse dummy is included; the values increase from 0.29 to 0.56 and from 0.10 to 0.48, respectively. None of the models suffer from serial correlation or ARCH effects. Including the contemporary and one-period lagged first difference (further lags are not significant) of the oil price in the short-run dynamics results in normally distributed residuals for models 3 and 3R. The oil price thus seems to compensate for supply shocks hitting the economy.

The real interest rate is always significant, and its coefficient remains stable for different specifications. The vacancy ratio becomes significant after correcting for an outlier in the short-run dynamics. Its coefficient also remains more or less stable for models 2, 3 and 3R. In the first two models, the NAIRU enters the long-run relationship significantly, with a rather high coefficient between -3.7 and -4.3. However, including the oil price leads to an insignificant NAIRU coefficient. In model 3R, we therefore

[7]It is plausible to assume that the oil price is not affected by economic conditions in Germany.

[8]We also attempt to normalise the NAIRU series, but as the adjustment coefficient for this equation is not significantly negative, it would lead to a misspecification.

Table 1: VECM estimation results

	Model 1	Model 2	Model 3	Model 3R
	T = 67	T = 67	T = 67	T = 67
Beta				
RLTIR	0.784 (0.182)	0.888 (0.198)	0.829 (0.153)	0.819 (0.113)
VACRATIO	0.120 (0.092)	0.228 (0.099)	0.314 (0.109)	0.300 (0.051)
LWS	1.000 (0.000)	1.000 (0.000)	1.000 (0.000)	1.000 (0.000)
NAIRU	-4.385 (1.899)	-3.793 (2.080)	0.107 (2.020)	0.000 (0.000)
OILPRICE	-	-	-1.320 (0.466)	-1.290 (0.347)
Alpha				
RLTIR	-0.050 (0.029)	-0.043 (0.025)	-0.044 (0.033)	-0.045 (0.033)
VACRATIO	-0.300 (0.067)	-0.236 (0.058)	-0.363 (0.079)	-0.357 (0.074)
LWS	-0.130 (0.041)	-0.141 (0.028)	-0.186 (0.040)	-0.193 (0.039)
NAIRU	0.010 (0.011)	0.004 (0.007)	0.005 (0.010)	0.000 (0.000)
	Eq.1/Eq.2/Eq.3/Eq.4	Eq.1/Eq.2/Eq.3/Eq.4	Eq.1/Eq.2/Eq.3/Eq.4	Eq.1/Eq.2/Eq.3/Eq.4
Adj. R^2	0.28/0.57/0.29/0.10	0.28/0.59/0.56/0.48	0.34/0.60/0.53/0.46	0.36/0.61/0.54/0.47
SC(2), p-value	0.99/0.24/0.78/0.83	0.81/0.28/0.11/0.22	0.94/0.58/0.22/0.26	0.94/0.58/0.24/0.26
SC(4), p-value	0.99/0.46/0.36/0.98	0.96/0.50/0.22/0.47	0.99/0.75/0.40/0.51	0.99/0.76/0.43/0.47
ARCH(2), p-value	0.30/0.90/0.67/0.77	0.29/0.56/0.26/0.59	0.28/0.52/0.26/0.58	0.28/0.50/0.27/0.51
ARCH(4), p-value	0.38/0.99/0.93/0.74	0.24/0.88/0.61/0.62	0.41/0.73/0.51/0.67	0.41/0.73/0.52/0.62
NORM, p-value	0.00	0.01	0.34	0.39

Notes: Standard errors are given in parentheses. SC(p), ARCH(p) and NORM refer, respectively, to the Ljung-Box Q test ($\chi^2(p)$) on serial correlation, the LM test ($\chi^2(p)$) on ARCH effects and the Doornik Hansen test (χ^2) on multivariate normality.

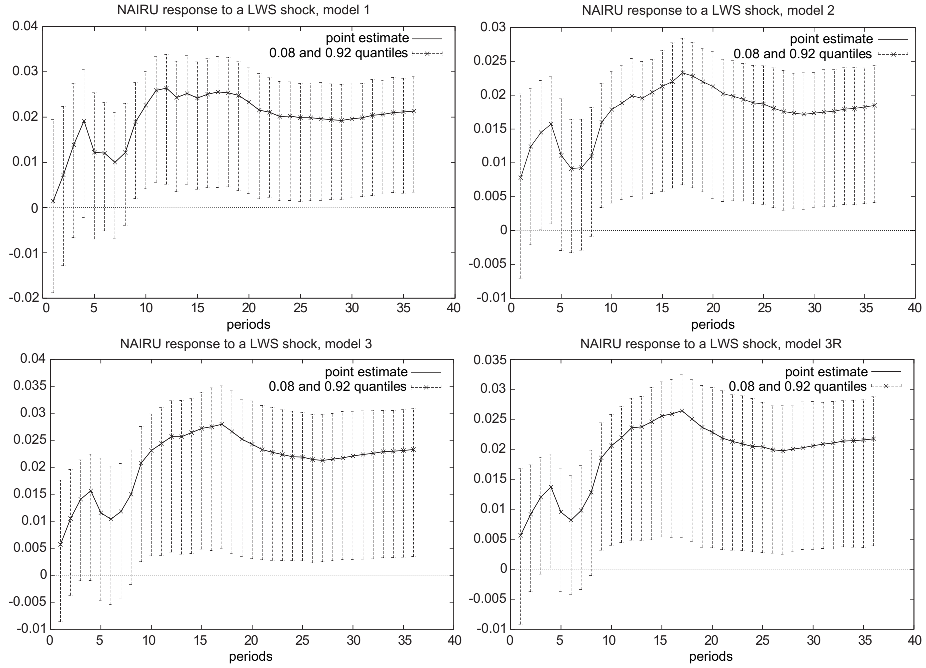

Figure 3: Impulse-response functions of NAIRU to a shock in LWS

estimate a restricted VECM. We cannot reject the null hypotheses that the NAIRU does not adjust to pre-period, long-run disequilibrium and that it does not enter the long-run relationship (beta[NAIRU] = 0, alpha[NAIRU] = 0, LR test p-value = 0.855). We suppose that the oil price accounts for some structural factors affecting the economy, which the NAIRU does not fully explain; the cost factor information contained in the oil price seems to dominate the structural issues reflected by the NAIRU.

As a next step, we compute the impulse-response functions for 36 periods to describe the response of the low-wage share to a NAIRU shock, and vice versa. We use an 84% confidence interval, on the basis of 999 bootstrap replications.[9] The contemporaneous (causal) recursive Cholesky ordering is as follows: real interest rate $\rightarrow$ vacancy ratio $\rightarrow$ low-wage share $\rightarrow$ NAIRU. That is, we assume that the real interest rate affects aggregate demand and supply conditions, which leads to a reconsideration of employment demand and therefore changes in the vacancy ratio. This shift affects the low-wage extent through the employment ratio between low-wage and non-low-wage workers. Finally, we assume the NAIRU takes the last position, because it is affected by several channels, as we discussed in Section 2. However, the impulse-response functions are robust to alternative orderings of these variables.

[9]With our small sample, we assume that the uncertainty of the estimates is relatively high. Therefore, we refrain from calculating more precise confidence intervals.

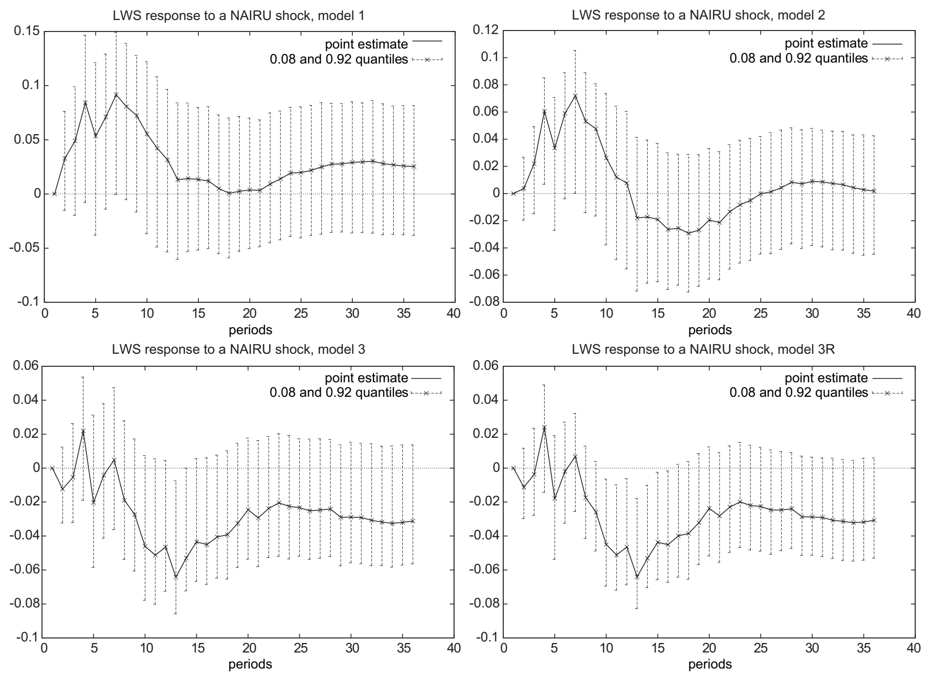

Figure 4: Impulse-response functions of LWS to a shock in NAIRU

Figure 3 depicts the impulse responses of the NAIRU to a positive shock in the low-wage share. For all four models, we find a significant positive response after around two years, in line with the findings of Garz (2010). Thus, the presumption that an increase in the low-wage extent results in a lower NAIRU, as in standard wage/price-setting theory, is strongly rejected. Instead, the results from this symmetric approach favour the Keynesian argument.

We depict the impulse-response functions for the reverse direction in Figure 4. None of the four models indicates a significant impact of a NAIRU shock on the low-wage share. The initial response of the point estimator tends to be positive for models 1 and 2 and approximately 0 for models 3 and 3R. In the medium to long run, the point estimator of model 1 remains positive, approximately 0 for model 2 and negative after the inclusion of the oil price (model 3) or the imposed restriction on the NAIRU (model 3R). Thus, the effect of an unexpected NAIRU change on the extent of the low-wage sector remains ambiguous.

5.2 Asymmetric error correction estimates

We employ the same variables we used for the VECM estimation. Because the asymmetric approach is limited to a single-equation representation, we estimate both directions, first assuming that the low-wage share, and then the NAIRU, is endogenous:

1. LWS = f(NAIRU, RLTIR, VACRATIO, OILPRICE),

2. NAIRU = f(LWS, RLTIR, VACRATIO, OILPRICE).

For each direction, we estimate three different specifications. First, we start by estimating the pure model. Second, we add impulse dummies to account for outliers. Third, we add the level of the oil price, as well as its contemporaneous and lagged first differences. The maximum lag length for models 1 and 2 is five, whereas the small sample size limits model 3 to four lags. We follow a general-to-specific approach to select the final lag structure by successively dropping all lags insignificant at the 5% level.

Table 2 shows the long-run estimation results, with low-wage extent as the dependent variable. For all three specifications, the PSS F-test and Banerjee et al (1998) BDM t-test suggest a significant long-run relationship. The fit of the models is remarkably high, with the lowest (adjusted) R^2 value around 0.67. Our battery of tests indicates some potential serial correlation for models 1 and 3 (the null hypothesis can only be rejected at the 5% level), some potential model misspecification for model 1 (RESET test) and some ARCH effects for model 3. The latter model also seems to be subject to parameter instability, as suggested by the QLR test around 1998Q2.

The estimated coefficients indicate that decreasing structural unemployment (i.e. increase in the NAIRU(-) variable) is accompanied by a smaller low-wage extent. In turn, increasing structural unemployment (i.e. NAIRU(+) increase) leads to a higher low-wage share. The estimated coefficient for the NAIRU(+) series is only significant in models 2 and 3. In contrast, the NAIRU(-) coefficient is highly significant in the first two models but slightly above 10% when we include the oil price. Two explanations could indicate why structural unemployment affects the low-wage sector as these estimates suggest. First, German labour market reforms, which contributed to the low-pay expansion, are clearly reactions to the high, long-lasting levels of unemployment. Second, a lower NAIRU likely increases the bargaining power of workers at the lower end of the wage distribution. These workers are usually less qualified, which allows firms to replace them easily. However, their disadvantageous bargaining position improves when unemployment decreases and more jobs become available. If they are able to bargain for higher real wages, some workers may exit the low-pay sector. In contrast, an increase in structural unemployment reduces the bargaining power of low-wage employees, which causes the low-pay sector to grow instead. Labour market institutions probably induce an asymmetric form in this effect, as the coefficients and the bootstrapped Wald test indicate. The lower absolute magnitude of the NAIRU(+) coefficient seems to reflect downward wage rigidity, caused by labour market regulation (e.g. employment protection legislation, minimum and combination wages) and unions, which jointly work to combat increasing inequality when structural downturns appear. In contrast, these institutions do not impede the reduction of wage inequality if employment rises permanently. This asymmetry can be visualised according to the dynamic multipliers in Figure 5 (panels a, c and e). These graphs also indicate that the shock reaction reaches its peak after a few periods and remains stable in the medium and long runs.

Table 3 depicts the other case, with the NAIRU as the dependent variable. For all three specifications, the PSS F-test suggests at least one long-run relationship. The BDM t-test indicates a cointegration relationship for models 1 and 3 but not for model

Table 2: AECM estimation results (dependent variable: d_LWS)

	Model 1, T = 66 Coefficient (p-val.)	Model 2, T = 66 Coefficient (p-val.)	Model 3, T = 67 Coefficient (p-val.)
L_RLTIR(+)	-0.086 (0.762)	-0.062 (0.799)	-0.368 (0.413)
L_RLTIR(-)	-0.253 (0.029)	-0.186 (0.091)	-0.141 (0.634)
L_VACRATIO(+)	-0.260 (0.000)	-0.259 (0.000)	-0.242 (0.010)
L_VACRATIO(-)	-0.107 (0.060)	-0.078 (0.165)	-0.190 (0.121)
L_NAIRU(+)	0.852 (0.154)	1.560 (0.008)	1.906 (0.089)
L_NAIRU(-)	-1.827 (0.025)	-1.723 (0.038)	-2.573 (0.111)
L_OILPRICE(+)			-0.312 (0.704)
L_OILPRICE(-)			0.242 (0.825)
Wsym_RLTIR, p-value	0.452	0.536	0.574
Wsym_VACRATIO, p-value	0.090	0.040	0.779
Wsym_NAIRU, p-value	0.015	0.003	0.046
Wsym_OILPRICE, p-value			0.709
Adj. R^2	0.713	0.699	0.674
SC(4), p-value	0.094	0.207	0.087
ARCH(4), p-value	0.819	0.123	0.062
NORM, p-value	0.734	0.367	0.945
RESET, p-value	0.080	0.141	0.224
CUSUM, p-value	0.643	0.578	0.490
QLR, p-value (break date)	>10% (2003Q3)	>10% (2003Q3)	<5% (1998Q2)
BDM, t-stat (t-crit 1/5/10%)	-10.11 (-4.99/-4.38/-4.04)	-9.17 (-4.99/-4.38/-4.04)	-5.53 (-5.37/-4.72/-4.40)
PSS, F-stat (F-crit 1/5/10%)	20.17 (4.43/3.61/3.23)	13.90 (4.43/3.61/3.23)	8.37 (4.10/3.39/3.06)

Notes: (+) and (-) denote positive and negative partial sum processes, respectively. Wsym_* refers to the bootstrapped Wald test statistic for long-run symmetry. SC(4), ARCH(4), NORM, RESET, CUSUM and QLR denote, respectively, the tests for serial correlation (Breusch-Godfrey, F-statistic), ARCH effects ($\chi^2(4)$), normality (χ^2), functional form (Ramsey's RESET, F-statistic), parameter stability (Harvey-Collier test statistics) and a structural break at an unknown point (15% trimming). BDM refers to the cointegration test suggested by Banerjee et al (1998), which provides t-statistics and corresponding critical values. PSS denotes the F-statistic and critical values for the bounds testing approach proposed by Pesaran et al (2001).

2. The fit of the first model is relatively low (adjusted $R^2 = 0.346$). It is not lower than the (adjusted) $R^2 = 0.738$ for the other two models, which is probably due to the inclusion of impulse dummies and the oil price. For model 1, the tests indicate some misspecification, potential ARCH effects and a structural break in 2005Q1. For model 3, we also find some evidence for a structural break in 2004Q4.

The LWS(+) coefficient, which measures the impact of a growing low-pay sector, is only significant in model 1, which indicates a positive impact on the NAIRU, in line with the VECM. In model 2, the LWS(+) coefficient is negative but not significant. Model 3 again shows the expected positive sign for this coefficient, though it is not significant. The VECM results receive further confirmation from the coefficient estimate for LWS(-), which is significantly negative in model 2 (a shrinking low-pay sector lowers structural unemployment). In model 1 and 3, the LWS(-) coefficient admittedly has a positive sign but is not significant. However, the VECM results are always confirmed when the LWS(+) and LWS(-) estimate is significant. The null hypothesis of symmetry cannot be rejected for any of the models, which is probably associated with the broadly indicated insignificance of the LWS(+/-) coefficients. For the same reason, the dynamic multipliers in Figure 5 (panels b, d and f) are difficult to interpret. They indicate a NAIRU response that is either positive or negative, irrespective of whether it is caused by a LWS(+) or LWS(-) shock.

6 Conclusion

Based on quarterly data from 1991Q1 to 2008Q4, we investigate the relationship between structural unemployment and the extent of the low-wage sector in Germany. Controlling for other structural and institutional influences, we conclude from the VECM that the causality in this context runs from the extent of the low-pay sector to structural unemployment. But contrary to standard wage/price-setting theory, the relationship is not negative. Instead, an increasing share of low-wage earners raises structural unemployment. The finding corresponds to the Keynesian perspective that a growing low-wage sector lowers consumption, which is not fully compensated for by reduced labour costs and the associated increase in investment demand. The resulting reduction of aggregate demand finally leads to higher structural unemployment. If they have any related effect, the creation of low-wage jobs therefore might contribute to reduce cyclical, but not structural, unemployment. This is problematic, because the German low-wage sector has been growing steadily since 1995 and is presumably expanding further as intended by labour market policy. The result of this study also implies that the hoped-for effect of getting people into stable, long-term employment through their initial experience with flexible relationships (the 'stepping stone' effect) does not seem very likely.

With the AECM we do not find one-sided causality, but we identify some reverse effects. Changes in structural unemployment have repercussions for the share of low-wage earners, because they influence the bargaining power of less qualified and easily replaceable workers. The AECM also indicates asymmetry in this effect. That is, institutions such as labour market regulations, unions and collective bargaining have a preserving

Table 3: AECM estimation results (dependent variable: d_NAIRU)

	Model 1, T = 66 Coefficient (p-val.)	Model 2, T = 66 Coefficient (p-val.)	Model 3, T = 66 Coefficient (p-val.)
L_RLTIR(+)	-0.128 (0.037)	-0.103 (0.039)	0.082 (0.299)
L_RLTIR(-)	0.014 (0.661)	-0.002 (0.175)	-0.208 (0.003)
L_VACRATIO(+)	-0.043 (0.000)	-0.117 (0.978)	-0.024 (0.176)
L_VACRATIO(-)	-0.060 (0.000)	0.009 (0.001)	-0.058 (0.008)
L_LWS(+)	0.095 (0.039)	-0.371 (0.775)	0.023 (0.618)
L_LWS(-)	0.017 (0.722)	-0.318 (0.000)	0.058 (0.308)
L_OILPRICE(+)			-0.294 (0.112)
L_OILPRICE(-)			0.982 (0.001)
Trend		0.036 (0.038)	
Wsym_RLTIR, p-value	0.048	0.454	0.007
Wsym_VACRATIO, p-value	0.320	0.017	0.325
Wsym_LWS, p-value	0.232	0.483	0.613
Wsym_OILPRICE, p-value			0.000
Adj. R^2	0.346	0.738	0.835
SC(4), p-value	0.950	0.885	0.631
ARCH(4), p-value	0.094	0.725	0.361
NORM, p-value	0.002	0.432	0.637
RESET, p-value	0.001	0.554	0.930
CUSUM, p-value	0.714	0.713	0.943
QLR, p-value (break date)	<1% (2005Q1)	>10% (1998Q2)	<5% (2004Q4)
BDM, t-stat (t-crit 1/5/10%)	-5.78 (-4.99/-4.38/-4.04)	-4.46 (-5.65/-5.01/-4.68)	-5.04 (-5.37/-4.72/-4.40)
PSS, F-stat (F-crit 1/5/10%)	5.36 (4.43/3.61/3.23)	7.04 (4.90/4.00/3.59)	7.18 (4.10/3.39/3.06)

Notes: (+) and (-) denote positive and negative partial sum processes, respectively. Wsym_* refers to the bootstrapped Wald test statistic for long-run symmetry. SC(4), ARCH(4), NORM, RESET, CUSUM and QLR denote, respectively, the tests for serial correlation (Breusch-Godfrey, F-statistic), ARCH effects ($\chi^2(4)$), normality (χ^2), functional form (Ramsey's RESET, F-statistic), parameter stability (Harvey-Collier test statistics) and a structural break at an unknown point (15% trimming). BDM refers to the cointegration test suggested by Banerjee et al (1998), which provides t-statistics and corresponding critical values. PSS denotes the F-statistic and critical values for the bounds testing approach proposed Pesaran et al (2001).

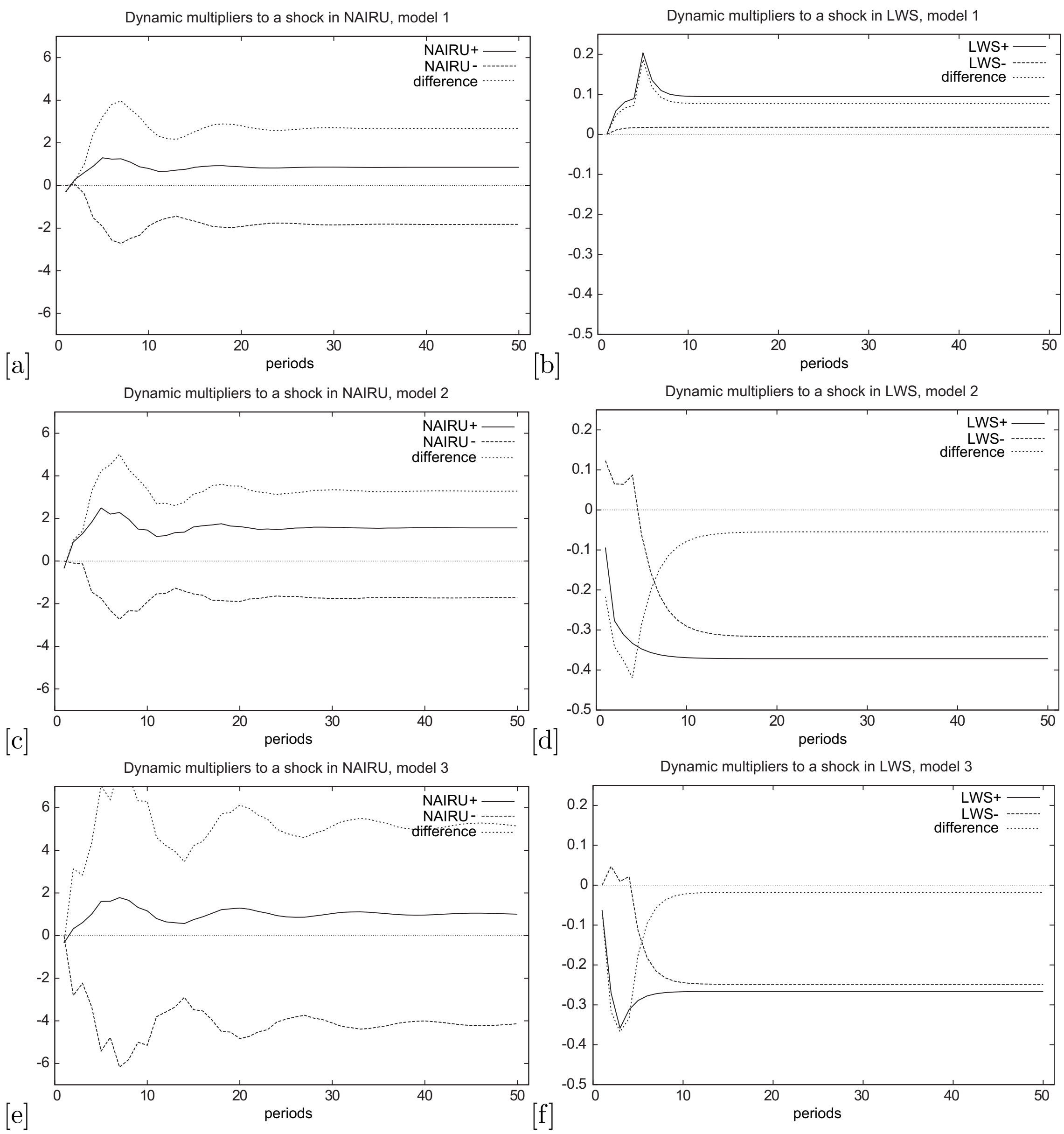

Figure 5: Dynamic multipliers

effect on wage equality in periods of structural downturns, but they do not interfere when the labour market tightens.

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A Estimation of the NAIRU

This section briefly describes the NAIRU estimation, as presented by Garz (2010), based on a theoretical derivation from the wage/price-setting system discussed by Layard et al (1991) and Turner et al (2001). The NAIRU is estimated in a state space system, where the observation equation represents the Phillips curve:

$$\Delta\pi_t = \alpha(L)\Delta\pi_{t-1} + \beta(L)Gap_t + \gamma(L)Pim_t + \delta D_t + \varepsilon_t^{\pi} \,. \tag{8}$$

The inflation rate is denoted as π_t, where Δ indicates the first difference. Lagged values of the change of inflation are allowed to enter as explanatory variables. Gap_t represents the unemployment gap, and Pim_t is the import price index. To capture German reunification and some other outliers, the vector D_t contains several dummy variables. The error term ε_t^{π} is assumed to be independently and normally distributed. The observation equation is complemented by the definition of the unemployment rate as the sum of the unemployment gap and the NAIRU:

$$U_t = Gap_t + NAIRU_t \,. \tag{9}$$

The stochastic properties of the unemployment gap and the NAIRU are described by the following state equations:

$$Gap_t = ar_1 Gap_{t-1} + ar_2 Gap_{t-2} + \varepsilon_t^{Gap} \,, \tag{10}$$

$$NAIRU_t = NAIRU_{t-1} + \theta I_{91q1,t} + \varepsilon_t^{NAIRU} \,. \tag{11}$$

In equation (10), the unemployment gap is modelled as an AR(2) process. With the condition $|ar_1 + ar_2| < 1$, the cyclical component of unemployment is a stationary, autocorrelated process with a sample mean of 0. In the second state equation, the NAIRU is defined as a random walk without drift. The impulse dummy $I_{91q1,t}$ is included to account for the reunification break. The error terms of both state equations are assumed to be normally distributed and mutually uncorrelated.

This state space set-up is similar to the specifications provided by Laubach (2001), Logeay and Tober (2006), Stephanides (2006), and Turner et al (2001). The attempt to model the NAIRU as a random walk with drift, as in some of these studies, produces rather unsatisfactory results (i.e. both the trend term and the trend state are statistically insignificant). In addition, the inclusion of a stochastic trend term in the NAIRU equation characterises the unemployment rate as I(2), which is not supported by unit root tests.

The data for inflation and unemployment come from the OECD. The inflation rate (LOG_PINFL) is defined as the first difference of the logarithm of the consumer price index. The unemployment rate (UREG) is the ratio of registered unemployed persons to the total civilian labour force. The short-term supply shock variable is the logarithm of the import price index (LOG_PIM). This variable is provided in the International Financial Statistics database of the IMF. All variables are available on a quarterly basis and at least from 1970Q1 to 2009Q4. They either had been seasonally adjusted by the provider, or we adjusted them using the Census X12 procedure. All series refer to unified Germany from 1991Q1 and to West Germany prior to this date. In accordance with ADF unit root tests (available on request), we assume all time series to be I(1).

The Phillips curve variables and their lagged values are chosen on the basis of the significance levels, guided by a preliminary stepwise OLS regression. Starting with the general model, we remove the most statistically insignificant variables step by step, until all remaining variables are at least significant at the 5% level. In this context, we substitute the unobservable NAIRU with the HP-filtered unemployment rate. The potential explanatory variables are eight lags of the dependent variable, the unemployment gap with four lags, the first difference of the unemployment rate, the first difference of the import price index with four lags and several impulse dummies to account for German reunification and outliers.

We use the resulting model to specify the Phillips curve equation in the state space model, as well as to obtain the starting coefficient values for the Kalman filter.[10] The initial state values we chose equal the first observation of the HP-filtered unemployment rate and the corresponding unemployment gap. The initial variance-covariance matrix is diagonal, with large, arbitrarily set values. This approach allows the optimisation process to converge quickly.

Table 4 shows the NAIRU estimation results. It reports the estimated parameters of the state space system, the standard deviations of the respective error terms and tests for residual autocorrelation and normality. All estimated coefficients are in accordance with economic theory. The impulse dummies in the Phillips curve are not significant, which conflicts with the preliminary OLS estimate that indicated high significance. All impulse dummies are nonetheless retained in the system, because they contribute to residual normality. The sum of the autoregressive coefficients in the unemployment gap equation is slightly below 1.[11] The Phillips curve residual tests are satisfactory. Neither null hypothesis, of normally distributed or serially uncorrelated residuals, can be rejected. The estimated time path of the Kalman smoothed NAIRU is generally similar to those of previous studies with estimates for Germany (Fitzenberger et al (2008); Logeay and Tober (2006); SVR (2007); Turner et al (2001)).

To evaluate the robustness of the results, we also estimated the system with different shock variables (oil price, labour productivity and price wedge), an inflation rate based on the GDP deflator and an alternative measure of the unemployment rate. A detailed description of these variables and the results are available on request; the results do not substantially differ from our proposed specification. We also checked the robustness of

[10] Kalman filter estimates usually depend on exactly provided initial coefficient values. Tentatively changing these values shows that the estimation results are sensitive to such changes. In most cases, inappropriate starting values produce implausible results or prevent the optimisation process from converging. The same outcome applies if variables are added or removed without readjusting the initial coefficient values.

[11] If the sum of the autoregressive coefficients were exactly 1, the unemployment gap would not be stationary. ADF tests on the estimated gap series show that the null hypothesis of non-stationarity can be rejected with a p-value of 0.0040.

Table 4: NAIRU estimation results

	Dependent Variable: π_t^{CPI}
Sample	1984Q1 - 2008Q4 (T = 100)
Log Likelihood	475.4027
State Equations	
Gap_{t-1}	1.6738 (0.0000)
Gap_{t-2}	-0.7007 (0.0082)
$I_{91q1,t}$	0.0329 (0.9932)
Observation Equation	
$const$	0.0027 (0.0011)
π_{t-1}^{CPI}	0.2559 (0.0020)
π_{t-2}^{CPI}	0.1347 (0.0985)
Gap_t	-0.0006 (0.0412)
ΔPim_t	0.1239 (0.0000)
ΔPim_{t-1}	-0.0626 (0.0002)
ΔPim_{t-3}	0.0350 (0.0083)
$I_{91q3,t}$	0.0116 (0.5605)
$I_{91q4,t}$	0.0112 (0.3571)
$I_{93q1,t}$	0.0155 (0.3229)
Standard Deviations	
σ^{Gap}	0.1747
σ^{NAIRU}	0.1895
σ^{π}	0.0017
SC(4)	0.8132 (0.5202)
NORM	1.1820 (0.5538)

Notes: SC(4) and NORM denote the LM test for serial correlation (Breusch-Godfrey, F-statistic) and normality (χ^2), respectively. The p-values for the coefficients and test statistics are in parentheses.

our VECMs and AECMs with the alternatively obtained NAIRUs and again did not find any substantial differences.

B Calculation of the low-wage extent

The data for the calculation of the extent of the low-wage sector are provided by the German Socio-Economic Panel (SOEP), version 2008, of the German Institute for Economic Research (DIW). We include only dependently employed people. Workers older than 64 years, the usual age of retirement from the labour force in Germany, are removed. The lower age limit is predefined by the SOEP, because its surveys only include people older than 18 years of age. The extent of the low-wage sector can be calculated for each year from 1984 to 2008, the time span for which the necessary data are completely available. The initial period from 1984 to 1994 includes only West Germany and the time from 1995 to 2008 refers to the reunified country. For statistical inference, the data are cross-sectionally weighted.

The OECD defines a low-wage threshold as two-thirds of the median wage (OECD (1996)). We adopt this widely used definition. The reference wage is calculated on both a net and an hourly basis, which allows for a comparison across employees, regardless of their individual taxation or working time. The extent of the low-wage sector is then defined as the share of workers with wages below this threshold among all employees.

Because all other variables in this study are measured on a quarterly basis, we must disaggregate the resulting low-pay time-series. To do so, we applied the procedure suggested by Chow and Lin (1971). We checked several instruments[12] and decided to use labour productivity without the manufacturing industry as our auxiliary variable, because it led to the most plausible result. In addition, we applied an exponential moving average (current weight = 0.6) to remove some irregular high frequency movements.

C Unit root and cointegration tests

[12] Labour productivity of the manufacturing industry without the building sector, labour productivity of the total economy (both from the DESTATIS database of the German Federal Statistical Office) and real GDP (Eurostat database of the statistical office of the European Union).

Table 5: ADF-GLS tests

Variable	T-statistics	P-value	Deterministic	Lag (Max. 6)
NAIRU	-1.127	0.237	Constant	1
LWS	-1.300	>10%	Constant + Trend	4
RLTIR	-2.066	>10%	Constant + Trend	1
VACRATIO	-1.869	0.059	Constant	5
OILPROCE	-0.943	>10%	Constant + Trend	5
d_NAIRU	-11.505	0.000	Constant	0
d_LWS	-0.917	0.319	Constant	6
d_RLTIR	-2.409	0.015	Constant	2
d_VACRATIO	-1.997	0.044	Constant	4
d_OILPRICE	-1.367	0.160	Constant	6

Notes: The initial maximum lag length is 6, and the actual lag order is obtained by testing down as follows: estimating the Dickey-Fuller regression with k lags of the dependent variables; if the last lag is significant, executing the test with order k, otherwise, let $k = k - 1$. If case the test with linear trends using GLS p-values is not applicable, critical values from Elliott et al (1996), Table 1, are included instead.

Table 6: KPSS tests

Variable	T-statistics	P-value	Deterministic
NAIRU	0.697	0.015	Constant
LWS	0.170	0.037	Constant + Trend
RLTIR	0.279	<1%	Constant + Trend
VACRATIO	0.361	0.096	Constant
OILPROCE	0.371	<1%	Constant + Trend
d_NAIRU	0.235	>10%	Constant
d_LWS	0.164	>10%	Constant
d_RLTIR	0.227	>10%	Constant
d_VACRATIO	0.350	>10%	Constant
d_OILPRICE	0.062	>10%	Constant

Notes: Truncation lag = 3. The critical values shown for the test statistic are based on the response surfaces estimated by Sephton (1995), which are more accurate for small samples than the values given in the original KPSS article. When the test statistic lies between the 10% and 1% critical values, p-values, obtained by linear interpolation, are provided.

Table 7: Johansen cointegration tests for models 1 and 2

H0	H1	95% cv	90% cv	Trace (Model 1)	Trace (Model 2)
$r = 0$	$r = 1$ / $r >= 1$	48.880	45.700	76.239	86.560
$r <= 1$	$r = 2$ / $r >= 2$	31.540	28.780	21.611	21.669
$r <= 2$	$r = 3$ / $r >= 3$	17.860	15.750	10.572	5.712
$r <= 3$	$r = 4$	8.070	6.500	2.160	1.055

Notes: The underlying VAR models are of order 5 and contain unrestricted intercept coefficients. The statistics refer to Johansen's log-likelihood-based trace statistics and are computed using observations for the period 1991Q1 to 2008Q4. The asymptotic critical values are those provides by Pesaran et al (2000).

Table 8: Johansen cointegration test for model 3

H0	H1	95% cv	90% cv	Trace
$r = 0$	$r = 1$ / $r >= 1$	58.630	45.700	98.680
$r <= 1$	$r = 2$ / $r >= 2$	38.930	35.880	26.381
$r <= 2$	$r = 3$ / $r >= 3$	23.320	20.750	8.3972
$r <= 3$	$r = 4$	11.470	9.530	2.2362

Notes: The underlying VAR model is of order 5 and contains an unrestricted intercept, with the oil price treated as an exogenous I(1) variable. The statistics refer to Johansen's log-likelihood-based trace statistics and are computed using observations for the period 1991Q1 to 2008Q4. The asymptotic critical values are those provided by Pesaran et al (2000).